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A Forty Year Assessment of Forecasting the Boat Race

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A forty year assessment of forecasting the Boat Race

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Abstract

We study the forecasting of the yearly outcome of the Boat Race between Cambridge and Oxford. We compare the relative performance of different dynamic models for forty years of forecasting. Each model is defined by a binary density conditional on a latent signal that is specified as a dynamic stochastic process with fixed predictors. The out-of-sample predictive ability of the models is compared between each other by using a variety of loss functions and predictive ability tests. We find that the model with its latent signal specified as an autoregressive process cannot be outperformed by the other specifications. This model is able to correctly forecast 30 out of 40 outcomes of the Boat Race.

Keywords: Binary time series, Predictive ability, Non-Gaussian state space model

1 Introduction

The Boat Race between the universities of Cambridge and Oxford is an annual rowing event that started in 1829 and is by now the most popular one day sports event in the world. Each year, a crew from the University of Cambridge rows against a crew from the University of Oxford. The rowing crews consist of eight rowers and one cox. The race takes place on the river Thames between Putney and Mortlake. We consider a class of models for the Boat Race outcomes and investigate the ability of the models to forecast the next outcome of the Boat Race. More specifically, for a variety of dynamic binary model specifications, we assess their ability to forecast the outcome of the race between 1971 and 2010 using information that is available to us just prior to the race. The relative out-of-sample forecasting performance of the models is compared for different loss functions and equal predictive ability tests; see Diebold & Mariano (1995), Diebold & Lopez (1996) and White (2000).

Forecasting the Boat Race is interesting from a number of viewpoints. First, bookmakers and gambling men may increase their expected profits by using the forecasts. Especially, if the forecasts from the dynamic models are significantly better than the forecasts resulting from simple procedures, such as flipping a coin or setting the forecast equal to the winner of the previous race. Second, from an econometric forecasting perspective, the historical outcomes of the Boat Race form an interesting binary time series. While the number of observations can be limited, the observations themselves can be spread over a long period of time and with many missing entries. Binary time series occur in many fields of research such as finance, criminology and computer science. The generation of accurate forecasts for events (such as winning a race, intervening in a market or closing a computer network) can be of vital importance for many instances.

We first investigate the in-sample performance of different binary models. The models are summarized by a binary density for the outcome of the race that is defined conditional on a signal that depends on an observable deterministic component and on a latent, stochastic and dynamic component. The deterministic component includes a constant and a vector of predictors, such as the average difference in weight between the boats and the winner of the toss. The latent stochastic component is modeled by a Gaussian dynamic processes such as random walk, autoregressive, autoregressive fractionally integrated and stochastic cycle processes. The parameters of the models are estimated by the method of Monte Carlo maximum likelihood. In particular, for the stochastic processes that imply a short-memory process we implement the simulation-based methods of Shephard & Pitt (1997), Durbin & Koopman (1997) and Jungbacker & Koopman (2007). These methods are based on the importance sampling technique. When the stochastic component is a long-memory process we alter the estimation method based on the procedures discussed in Mesters, Koopman & Ooms (2011).

We further perform an out-of-sample forecasting study. Each model is used to compute a probability forecast for the event that Cambridge wins the next Boat Race. This is done repetitively for a period of forty years (1971-2010), where we only use information that would have been available just before the race starts. Forecasting in this manner is referred to as “pseudo-out-of-sample” forecasting by Stock & Watson (2003). The relative out-of-sample predictive performance of the different models is compared by different loss functions, or score statistics. We verify whether the models possess significantly different predictive abilities as implied by different loss functions, see Diebold & Mariano (1995) and West (1996).

The remainder of our paper is organized as follows. We continue the introduction with a short outline of the history of the Boat Race and describe the data that is obtained from the history of the race. Section 2 discusses the binary models that we consider for forecasting the race. Section 3 discusses the forecasting methodology and the predictive ability tests. It further presents and discusses the results of forty years of out-of-sample forecasting. Section 4 provides a brief summary of our study.

1.1 History

The first race between the universities of Oxford and Cambridge was organized in 1829. The idea came from two friends; Charles Merivale, a student at Cambridge, and Charles Wordsworth, a student at Oxford. On 12 March 1829 Cambridge sent a challenge to Oxford and the history of the race started. The first race was at Henley-on-Thames and was won by Oxford. In 1839 the race was relocated to London, by now the race had become an annual event taking place between Westminster and Putney. However, the increased crowds interested in the race made it necessary to move yet again. In 1845 the race was first held on the course between Putney and Mortlake, which is also on the river Thames. In 1836 the Oxford crew selected the color dark blue to race in and the Cambridge crew selected the “duck egg blue” of Eton to race in.

The race outcome were about even until 1861, when Oxford started the first winning streak in the history lasting nine years. In 1877, the race was declared a dead heat for the first and only time in its history - although legend has it that the judge, “Honest John” Phelps, was asleep under a bush when the crews reached the finishing line. The 1912 event witnessed another Boat Race first when both boats sank and the race had to be re-run the next day, with Oxford claiming the honors at the second attempt. The race was not held between 1915 and 1919 due to the First World War. When it resumed in 1920, Cambridge embarked on a lengthy period of domination. They would win the race 13 years running between 1924 and 1936, the longest winning streak in the race’s history.

There was another break between 1940 and 1945 because of the Second World War, although four unofficial races were held during this time, all outside London. The 1952 contest witnessed perhaps the most extreme weather in Boat Race history, with Oxford prevailing in the midst of a blizzard. The dark blues also won the 100th Boat Race in 1954. In 1981, Sue Brown became the first female to enter the Boat Race, acting as cox for Oxford. The following year, Hugh and Rob Clay of Oxford became the first twins to win the race. The dark blues dominated throughout the Eighties, as Cambridge suffered a series of misfortunes. The biggest of these came in 1984, when they managed to write off their boat before the start of the race. Controversy engulfed Oxford at the 1987 race when a section of the crew rose up in mutiny against the president over team selection policy. However, the dispute, which was chronicled in the book and film ‘True Blue’, did not prevent them from winning the race again.

Cambridge regained their pride in 1993 by ending Oxford’s domination. They subsequently won the race seven years running, the highlight coming in 1998 when they broke the course record by a massive 26 seconds. The last race that we consider was won by Cambridge in 2010.

1.2 Data

The dependent variables in our study are denoted by the elements of the $n \times 1$ vector $\mathbf{y} = (y_1, \dots, y_n)'$, which indicate the outcome path of the Boat Race. All data is freely

obtained from <http://www.theboatrace.org>. For each year t , where $t = 1$ refers to the year 1829 and $t = n$ to the year 2010, we let $y_t = 1$ indicate that Cambridge has won the race in year t , whereas $y_t = 0$ indicates that Oxford has won the race in year t . For some years the race was not held due to a variety of circumstances, for these years we consider the corresponding outcome missing. In particular, the years 1830-1825, 1837, 1838, 1843, 1844, 1847, 1848, 1850, 1851, 1853, 1855, 1877 (dead heat), 1915-1919 (WWI) and 1940-1945 (WWII) are considered missing. In total we have $n = 182$ observations, of which 28 are missing. Cambridge currently leads the series, with 80 wins compared to 74 wins for Oxford.

During the history of the race the average weight of the rowers has increased substantially. This has led to the suspicion that increasing the weight of the rowers leads to faster boats, due to the presents of more muscle power. The contrary could however also be possible, by reasoning that a lighter boat has less water resistance and would therefore be faster. We include the average log difference in weight between the rowers in the Cambridge and Oxford boats as the first predictor.

The second time-varying predictor that we include is the outcome of the coin toss. The club presidents toss a coin before the race for the right to choose which side of the river (station) they will row on. Their decision is based on the weather conditions and how the various bends in the course might favor the pace of their crew. The north station (Middlesex) has the advantage of the first and last bends, whereas the south (Surrey) station has the longer middle bend. It is generally believed that the winner of the coin toss has an improved chance of winning. In most years, the betting odds change severely after the toss outcome has been made public.

The $n \times 2$ matrix \mathbf{X} includes the time-varying predictors, where element $x_{t,1}$ is the average log difference in weight between the rowers of Cambridge and Oxford in year t and $x_{t,2} = 1$ if Cambridge wins the toss in year t and zero otherwise.

2 Models and parameter estimation

The models that we use as our basis to forecast the outcome path of the race can be split into two parts; a conditional binary density for the observations and a signal that includes all dynamics and predictors. The conditional binary density is given by

$$p(y_t|\pi_t) = \pi_t^{y_t}(1 - \pi_t)^{1-y_t}, \quad t = 1, \dots, n, \quad (1)$$

where π_t denotes the winning probability for Cambridge in year t . Time-varying probability π_t is unknown to us, but is clearly restricted between zero and one. Therefore we specify the transformed probability as $\theta_t = \log(\pi_t/(1 - \pi_t))$ where the transformation is performed by the canonical link function for binary models; see Cox & Snell (1989). We refer to θ_t to as the signal for year t . The conditional density for the observations, given in (1) can be

rewritten in terms of the latent signal as

$$p(y_t|\theta_t) = \exp[y_t\theta_t - \log(1 + \exp \theta_t)], \quad t = 1, \dots, n, \quad (2)$$

which is assumed independent over $t = 1, \dots, n$. It holds that $p(\mathbf{y}|\boldsymbol{\theta}) = \prod_{t=1}^n p(y_t|\theta_t)$, where $\boldsymbol{\theta} = (\theta_1, \dots, \theta_n)'$. We decompose the signal in an observed deterministic part and a latent stochastic part. The deterministic part includes the intercept and two predictors (difference in weight of boat and winner of toss). The latent stochastic part captures the possible dynamic effects which reflect that the events of previous years also affect outcome of the race this year. The signal decomposition is given by the regression equation

$$\theta_t = \mu + \mathbf{x}_t\boldsymbol{\beta} + u_t, \quad t = 1, \dots, n, \quad (3)$$

where μ is the intercept, the 2×1 parameter vector $\boldsymbol{\beta}$ measures the effect of the observable predictors \mathbf{x}_t and u_t is the latent stochastic part. The 1×2 vector \mathbf{x}_t is the t th row of matrix \mathbf{X} . The models that we consider differ in the way that the latent process u_t is specified. Possible specifications that we include range from setting it to zero, which leads to a logistic regression model, to a complete dynamic specification. In the next subsection we discuss all specifications in some detail.

2.1 Signal specifications

Many dynamic specifications for u_t are possible. In general, we can expect to have a serial correlation structure between outcomes in successive years due to overlapping crews and coaches. For example, during the first winning streak of Oxford, between 1862 and 1869, the crew had the same coach George Morrison for six years who was the former oarsman of the crew. Furthermore, technological developments resulting from training methods or boat construction may provide comparative advantages to a team that can last for a number years. We focus on models for u_t that are able to pick up this kind of dynamic effects.

- (i) The simplest possible specification is the “constant” specification given by

$$u_t = 0 \quad t = 1, \dots, n, \quad (4)$$

which results in a signal that is completely determined by the intercept and the two observed predictors. This is the baseline specification, which assumes that the entire signal is observable. In effect the model for y_t is reduced to a logistic regression model.

- (ii) The first specification that we consider to capture serial correlation for u_t is the random walk as given by

$$u_{t+1} = u_t + \eta_{t+1}, \quad \eta_{t+1} \sim N(0, \sigma_\eta^2), \quad t = 2, \dots, n, \quad (5)$$

where the next value of u_t is equal to the old value plus an error term. The initial

state, u_1 , has an unknown distribution and is therefore fixed at zero. This corresponds to a fifty percent winning probability for Cambridge without covariates.

- (iii) The dependence of the outcome on previous realisations of u_t can be tampered by using an autoregressive (AR) specification. The coefficients of the $\text{AR}(p)$ polynomial $\phi(B) = 1 - \phi_1 B - \dots - \phi_p B^p$ enable the adjustment of observations from the past, where B denotes the back shift operator as defined by $B^s y_t = y_{t-s}$ for any integer $s > 0$. The latent process u_t is given by

$$\phi(B)u_{t+1} = \eta_{t+1}, \quad \eta_{t+1} \sim N(0, \sigma_\eta^2), \quad t = p+1, \dots, n. \quad (6)$$

When we require to impose a stationary process for u_t , as opposed to the non-stationary random walk process (5), we need to assume that the roots of the autoregressive polynomial $\phi(B)$ lie outside the unit circle. The process is initialized for u_1, \dots, u_p by the unconditional distribution. For example, when $p = 1$ we have

$$u_1 \sim N\{0, \sigma_\eta^2/(1 - \phi_1^2)\}.$$

- (iv) Since some of the winning streaks are relatively long, we also consider a long memory specification for the latent process u_t . In particular, we investigate the autoregressive fractionally integrated (ARFI) model; see Granger & Joyeux (1980) and Palma (2007). The model for u_t can be expressed by

$$\phi(B)(1 - B)^d u_{t+1} = \eta_{t+1}, \quad \eta_{t+1} \sim N(0, \sigma_\eta^2), \quad t = p+1, \dots, n, \quad (7)$$

where the fractional integration part can be expressed by the binomial expansion

$$(1 - B)^d = \sum_{k=0}^{\infty} \frac{\Gamma(d+1)}{\Gamma(k+1)\Gamma(d-k+1)} (-1)^k B^k.$$

The fractional parameter d is a real valued constant in the range of $-1 < d < 0.5$. This assumption together with the previous assumption on the autoregressive polynomial ensures that the process is stationary, invertible and causal; see Palma (2007) for a proof.

- (v) Another possible component that may be present in some time series is the stochastic cycle process of Harvey (1989). To have a stochastically time-varying cycle component for u_t , we consider the following specification

$$\begin{bmatrix} u_{t+1} \\ u_{t+1}^* \end{bmatrix} = \phi \begin{bmatrix} \cos \lambda & \sin \lambda \\ -\sin \lambda & \cos \lambda \end{bmatrix} \begin{bmatrix} u_t \\ u_t^* \end{bmatrix} + \begin{bmatrix} \eta_{t+1} \\ \eta_{t+1}^* \end{bmatrix}, \quad t = 2, \dots, n, \quad (8)$$

where the errors term η_t and η_t^* are serially and mutually uncorrelated with mean zero and variance σ_η^2 . The parameter λ gives the frequency of the cycle and is estimated

together with the other parameters. The period of the cycle is $2\pi/\lambda$. Restrictions on the damping term, $0 \leq \phi \leq 1$, ensure that the process u_t is stationary. The cycle disturbance variance for both η_t and η_t^* is specified as $\sigma_\eta^2 = (1 - \phi^2)\sigma_u^2$ where the cycle variance σ_u^2 is for u_t and is treated as the parameter to estimate. The initial distribution of the cycle is then given by

$$u_1 \sim N(0, \sigma_u^2), \quad u_1^* \sim N(0, \sigma_u^2), \quad E(u_1 u_1^*) = 0.$$

It follows that a constant but stochastic cycle is obtained with $\phi = 1$ which remains properly defined; see Harvey & Koopman (1997).

The models that we consider can be summarized by the observational density (2), the signal (3) and one of the specification for u_t in (4), (5), (6), (7) and (8). For each model specification, the unknown parameters are collected in the vector $\boldsymbol{\psi}$.

2.2 Parameter estimation

Different estimation procedures may be considered for different model specifications. For exposition and comparison purposes we use the same Monte Carlo maximum likelihood estimation procedure for the estimation of the parameter vector $\boldsymbol{\psi}$ in all cases. In particular, we adopt an estimation procedure that is based on the simulation methods developed in Shephard & Pitt (1997) and Durbin & Koopman (1997).

The loglikelihood, for all models, is defined as $\ell(\boldsymbol{\psi}) = \log p(\mathbf{y}; \boldsymbol{\psi})$ where $p(\mathbf{y}; \boldsymbol{\psi})$ is the joint density of all observations. The terms after the semi-colon in $p(\cdot; \cdot)$ are the fixed, possibly unknown, arguments of the density function $p(\cdot; \cdot)$. The joint density for models defined by density (2) and signal (3) is given by

$$p(\mathbf{y}; \boldsymbol{\psi}) = \int_{\boldsymbol{\theta}} p(\mathbf{y}, \boldsymbol{\theta}; \boldsymbol{\psi}) d\boldsymbol{\theta} = \int_{\mathbf{u}} p(\mathbf{y}, \mathbf{u}; \mathbf{X}, \boldsymbol{\psi}) d\mathbf{u} = \int_{\mathbf{u}} p(\mathbf{y}|\mathbf{u}; \mathbf{X}, \boldsymbol{\psi}) p(\mathbf{u}; \boldsymbol{\psi}) d\mathbf{u}, \quad (9)$$

where $\mathbf{u} = (u_1, \dots, u_n)'$ and $p(\cdot)$, $p(\cdot, \cdot)$ and $p(\cdot|\cdot)$ denote marginal, joint and conditional density functions, respectively. Since \mathbf{X} and $\boldsymbol{\psi}$ are fixed we only need to integrate the stochastic process u_t from the joint density $p(\mathbf{y}, \mathbf{u}; \mathbf{X}, \boldsymbol{\psi})$. We further have

$$p(\mathbf{y}|\mathbf{u}; \mathbf{X}, \boldsymbol{\psi}) = p(\mathbf{y}|\boldsymbol{\theta}; \boldsymbol{\psi}) = \prod_{t=1}^n p(y_t|\theta_t; \boldsymbol{\psi}),$$

and

$$p(\mathbf{u}; \boldsymbol{\psi}) = p(u_1, \dots, u_p; \boldsymbol{\psi}) \prod_{t=p+1}^n p(u_t|u_1, \dots, u_{t-1}; \boldsymbol{\psi}).$$

A closed form solution for integral (9) is not available if u_t is stochastic and if $p(y_t|\theta_t; \boldsymbol{\psi})$ is binary. Instead we rely on numerical methods to solve the integral. In particular, we adopt the importance sampling method as it is discussed in, for example, Ripley (1987). It holds

that

$$p(\mathbf{y}; \boldsymbol{\psi}) = \int_{\mathbf{u}} \frac{p(\mathbf{y}|\mathbf{u}; \mathbf{X}, \boldsymbol{\psi})p(\mathbf{u}; \boldsymbol{\psi})}{g(\mathbf{u}|\mathbf{y})} g(\mathbf{u}|\mathbf{y}) d\mathbf{u} = g(\mathbf{y}) \int_{\mathbf{u}} \frac{p(\mathbf{y}|\mathbf{u}; \mathbf{X}, \boldsymbol{\psi})p(\mathbf{u}; \boldsymbol{\psi})}{g(\mathbf{y}|\mathbf{u})g(\mathbf{u})} g(\mathbf{u}|\mathbf{y}) d\mathbf{u}, \quad (10)$$

where $g(\mathbf{u}|\mathbf{y}) = g(\mathbf{y}|\mathbf{u})g(\mathbf{u})/g(\mathbf{y})$ is referred to as the importance density, which must be proportional to $p(\mathbf{y}, \mathbf{u}; \mathbf{X}, \boldsymbol{\psi})$, easy to sample from and easy to compute. In our approach, we take the importance density as Gaussian with a linear relation between y_t and $\theta_t = \mu + \mathbf{x}_t\boldsymbol{\beta} + u_t$ for $t = 1, \dots, n$ and for which it holds that $g(\mathbf{y}|\mathbf{u}) = \prod_{t=1}^n g(y_t|u_t)$. The mean and variance of the importance density is chosen such that the corresponding first two moments of $p(y_t|\theta_t; \boldsymbol{\psi})$ and $g(y_t|u_t)$ with respect to u_t are set equal. The resulting set of nonlinear equations can be solved iteratively using Kalman filter and smoothing methods; see Durbin & Koopman (2012, Section 10.6). A Monte Carlo estimator for (10) is given by

$$\hat{p}(\mathbf{y}; \boldsymbol{\psi}) = g(\mathbf{y}) M^{-1} \sum_{i=1}^M \frac{p(\mathbf{y}|\mathbf{u}^{(i)}; \mathbf{X}, \boldsymbol{\psi})p(\mathbf{u}^{(i)}; \boldsymbol{\psi})}{g(\mathbf{y}|\mathbf{u}^{(i)})g(\mathbf{u}^{(i)})}, \quad (11)$$

where samples $\mathbf{u}^{(i)}$ are drawn from $g(\mathbf{u}|\mathbf{y})$ for $i = 1, \dots, M$ using simulation smoothing methods.

For any importance density, it holds that $\hat{p}(\mathbf{y}; \boldsymbol{\psi}) \rightarrow p(\mathbf{y}; \boldsymbol{\psi})$ as $M \rightarrow \infty$, which is implied by the law of large numbers. The Lindeberg-Levy central limit theorem implies a \sqrt{M} convergence rate if draws from the importance sampler are independent and if importance weights

$$w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \boldsymbol{\psi}) = \frac{p(\mathbf{y}|\mathbf{u}^{(i)}; \mathbf{X}, \boldsymbol{\psi})p(\mathbf{u}^{(i)}; \boldsymbol{\psi})}{g(\mathbf{y}|\mathbf{u}^{(i)})g(\mathbf{u}^{(i)})} \quad i = 1, \dots, M. \quad (12)$$

have finite mean and variance; see the discussions in Geweke (1989). Failure of this condition to hold can lead to slow and unstable convergence of the estimator.

The details of the importance sampling methods for this class of time series models are discussed in Jungbacker & Koopman (2007). All our estimation results are based on this implementation. Mesters et al. (2011) have extended their procedures to allow for the long-memory (ARFI) specification of the signal in (7).

2.3 In-sample estimation results

Our sample of observations \mathbf{y} , the Boat Race outcome from 1829 until 2010, is used to optimize the simulated loglikelihood function based on (11) but subject to a bias-correction for the log transformation. For each model, we have taken $M = 1000$ draws from the constructed importance density and make use of three antithetic variables to relocate and rescale the weights. The simulated loglikelihood function is evaluated for different values of $\boldsymbol{\psi}$ during the numerical optimisation by using the same set of random input values to obtain a smooth loglikelihood function in $\boldsymbol{\psi}$. Most of such standard implementation details are discussed in Durbin & Koopman (1997).

In our discussion we distinguish models that include the two predictors in \mathbf{x}_t and models

that depend only on the dynamics in u_t . It allows us to assess whether including dynamic structures, such as the autoregressive structure, improves the model fit when compared to only having the intercept in the model. We include the following model specifications from Section 2.1 for the latent stochastic component u_t : constant, random walk, AR(1), ARFI(0, d), ARFI(1, d) and the stochastic cycle.

The parameter estimates of ψ for each specification are presented in Table 1. The top panel shows the results for models that include the predictors in \mathbf{x}_t . For all models, the estimated long-run mean μ is not significantly different from zero. Hence we can conclude that the models do not significantly predict Cambridge or Oxford as the winner for the Boat Race in the long run. The coefficient estimates for μ are all positive. This finding corresponds to the current lead in the series for Cambridge (80 wins against 74 for Oxford).

The difference in weight between the two boats is a strong predictor. The estimated corresponding coefficient is significant for all model specifications. Hence we may conclude that having more muscle power in the boat is more important than having less water resistance. Ideally, we would like to investigate this weight difference in more detail to verify whether the muscle power causes this effect or whether other related variables interact with the difference in weight. This requires additional information that is yet unavailable to us.

The outcome of the toss does not come forward as an important predictor. It is insignificant in all models with small coefficient values. This is rather surprising, given that the gambling odds usually change after the toss is made. It may have some impact when we account for certain weather conditions under which it may signify some advantage when a team starts at a specific side of the river.

The parameters that account for the dynamic effects in the model, including autoregressive coefficients, improve the model fit. For example, the parameter ϕ_1 (or ϕ) is estimated between 0.68 and 0.85 for the AR(1), ARFI(1, d) and stochastic cycle models. The long-memory parameter d is only positive and significant for the ARFI(0, d) model but it cancels out when the ϕ_1 parameter is also included as it is the case for the ARFI(1, d) model.

The stochastic cycle model shows overall the best in-sample fit according to the estimated loglikelihood value and the AIC criteria. The cycle frequency is estimated at 0.3338, which indicates a cycle period of nearly 18 years. This is a long cycle period and is mainly caused by the long winning streaks in the sample.

The parameter estimates for the model that do not include predictors are shown in the bottom panel of Table 1. We have also removed the insignificant mean μ for most of the models. The results are very similar when we compare these for models including predictors. Overall they perform less well in terms of likelihood and AIC values. The estimated dynamic parameters are however similar.

Figure 1 presents the estimated signal θ_t . The top panel shows them for models that include covariates. Here the estimated signals are mainly determined by the log difference in weight covariate. We experimented with different transformations for this variable, but the jumps in the weight difference remained large. The bottom panel for models without predictors presents smoother signal estimates.

	μ	ϕ	σ_η	λ	d	Including Predictors		β_1	β_2	$\log \hat{L}(\psi, \mathbf{y})$	AIC
Constant	0.135 _{0.241}						0.226 _{0.063}	0.168 _{0.350}	-239.872	485.745	
RW	0.396 _{0.850}		0.161 _{0.160}				0.221 _{0.066}	0.092 _{0.382}	-240.452	488.904	
AR(1)	0.321 _{0.437}	0.724 _{0.119}	1.064 _{0.404}				0.252 _{0.087}	-0.113 _{0.483}	-235.073	480.145	
ARFI(0, d ,0)	0.321 _{0.589}		1.319 _{0.502}		0.345 _{0.127}		0.270 _{0.091}	0.012 _{0.486}	-237.698	485.397	
ARFI(1, d ,0)	0.315 _{0.412}	0.762 _{0.178}	1.107 _{0.438}		-0.075 _{0.284}		0.254 _{0.089}	-0.108 _{0.487}	-235.037	482.073	
Cycle	0.288 _{0.365}	0.865 _{0.075}	0.755 _{0.317}	0.337 _{0.104}			0.244 _{0.0841}	-0.142 _{0.472}	-233.914	479.828	
Only Dynamics											
Constant	0.078 _{0.161}								-248.144	498.289	
RW			0.386 _{0.295}						-246.894	495.787	
AR(1)		0.753 _{0.101}	1.075 _{0.368}						-240.626	485.253	
ARFI(0, d ,0)			1.413 _{0.454}		0.391 _{0.104}				-243.803	491.605	
ARFI(1, d ,0)		0.771 _{0.159}	1.100 _{0.407}		-0.038 _{0.266}				-240.616	487.232	
Cycle		0.872 _{0.066}	0.764 _{0.288}	0.294 _{0.086}					-239.391	484.781	

Table 1: Parameter estimation results for the Boat Race from 1829 until 2010 ($n = 182$). The standard errors of the estimates are given in small print. The method of Monte Carlo maximum likelihood is based on importance sampling. The number of importance simulations for likelihood evaluation is $M = 1000$. The regression parameters β_1 and β_2 correspond to the effect of the average difference in weight between the rowers of the boats of Cambridge and Oxford (measured in 10 kilo units) and the effect of Cambridge winning the toss, respectively. The Akaike information criteria is computed as $2\mathbf{k} - 2 \log \hat{p}(\mathbf{y}; \hat{\psi})$, where \mathbf{k} denotes the number of parameters present in model k .

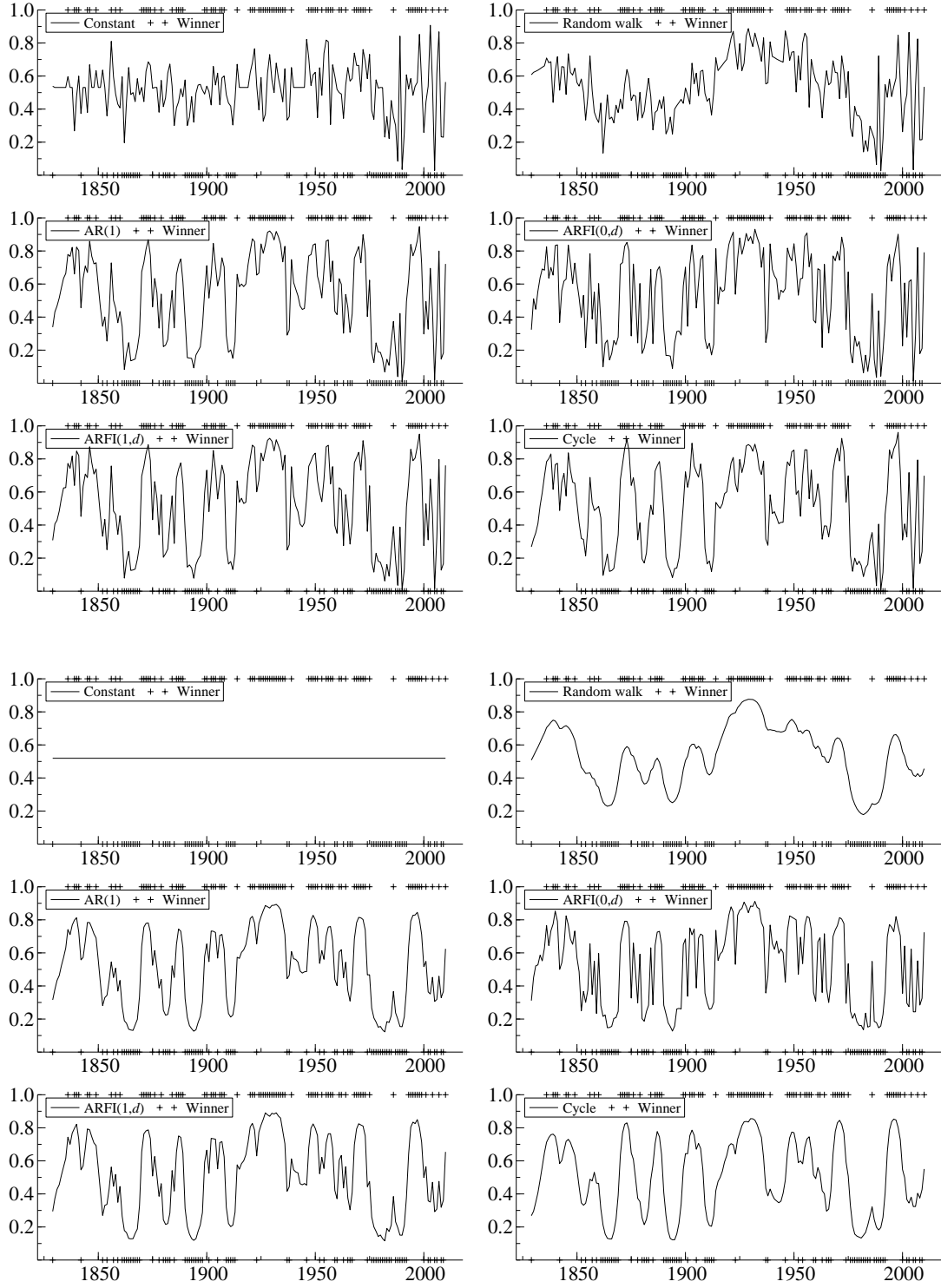


Figure 1: Estimated signals, $E_p(\theta|y; \mathbf{X}, \hat{\psi})$, for all models from Table 1. The top panel shows the signals for models including predictors \mathbf{X} and the bottom panel shows the signals for models that only include dynamics.

Model diagnostics

To guarantee a \sqrt{M} convergence rate for the Monte Carlo likelihood estimator (11), we must ensure that the weight function $w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \boldsymbol{\psi})$ in (12) has a finite variance as argued by Geweke (1989). To verify whether this is the case for our model specifications, we have implemented the Wald test as discussed by Koopman, Shephard & Creal (2009) and based on extreme value theory. The test statistic is evaluated at the estimated parameters in Table 1 and is constructed by the following step-wise procedure:

- (i) Given a model and its corresponding estimated parameters $\hat{\boldsymbol{\psi}}$, we generate 100,000 weight values $w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \hat{\boldsymbol{\psi}})$;
- (ii) Given a threshold value w^{\min} , we only consider weights that surpass the threshold;
- (iii) The resulting s exceedences z_1, \dots, z_s are modelled by the generalized Pareto density $f(w_i; a, b) = -\log b - (1 + a^{-1}) \log(1 + ab^{-1}z_i)$, for $i = 1, \dots, s$; the coefficients a and b are estimated by maximum likelihood straightforwardly; the estimates are denoted by \hat{a} and \hat{b} , respectively. The null hypothesis of variance existence is given by $H_0 : a \leq 0.5$.
- (iv) The test statistic $t_w = \hat{b}^{-1} \sqrt{s/3}(\hat{a} - 0.5)$ is under the null hypothesis asymptotically standard normal. We reject the null for a sufficiently positive value of t_w .

In Figure 2 we present the test statistics for the models in the top panel of Table 1. The test statistics are computed for a multiple of thresholds that result in sets of exceedence weights (with numbers of 1% to 50% of all weights). The test statistics are never rejected which indicate that the variance exists for many different threshold values. A similar panel of graphics can be produced for models without predictors.

Forecasting outlook

The likelihood values and the AIC criteria in Table 1 imply that the in-sample performance of our range of models is quite similar. It does not imply that the out-of-sample forecasting properties are also similar. We can illustrate this by presenting in Figure 3 the theoretical autocorrelation functions (ACF) of the estimated latent processes u_t . The ACFs are implied by the model specification with the parameters replaced by their estimates as given in the top panel of Table 1.

The constant model takes no serial correlation into account; its ACF is zero for all lags. All past outcomes of the Boat Race have an equally weighted impact on the forecast of the outcome this year. The random walk model is a highly persistent dynamic process resulting in a slowly decaying ACF. Most past outcomes have an impact on the current forecast but the most recent outcomes have a slightly higher weight. The AR(1) and ARFI(1, d) models have nearly identical autocorrelation functions. This is not surprising since the d and ϕ parameters in the ARFI(1, d) cancel out against each other to yield the short-memory AR(1) model. In both cases, only the more recent outcomes have an impact on the current forecast. The ACF of the ARFI(0, d) model shows hyperbolic decay, which is the characteristic of the

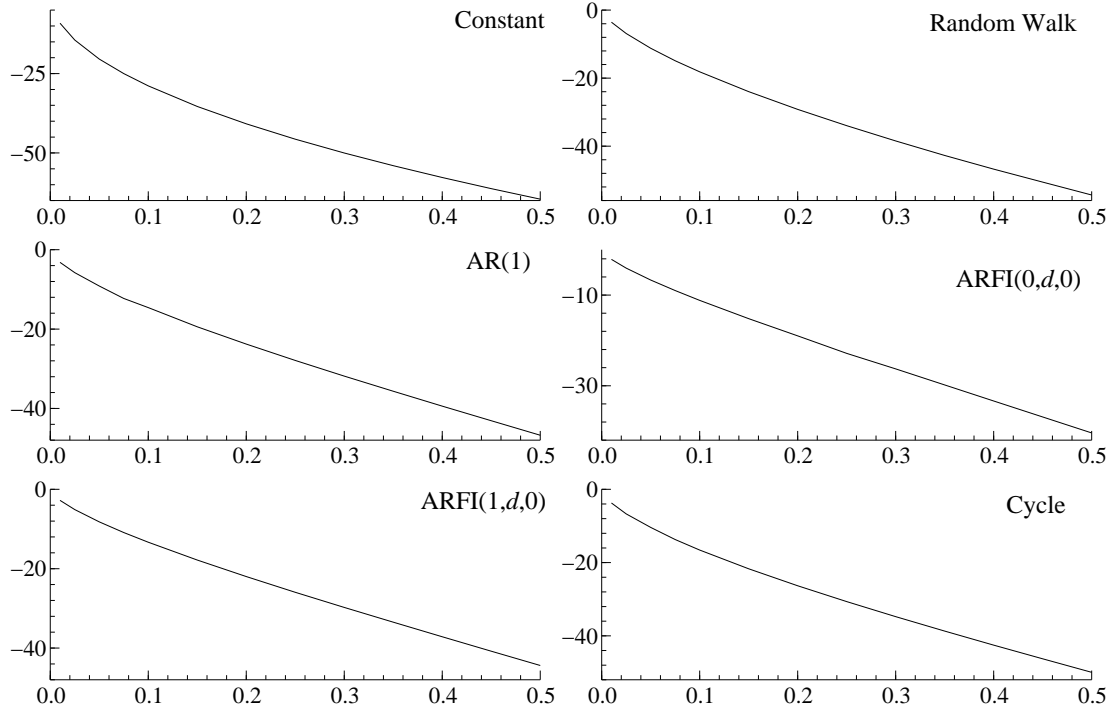


Figure 2: Importance sampling test statistics t_w for all models in the top panel of Table 1. The test statistics are based on 100,000 simulated weights $w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \psi)$. For each model we compute the test statistics for different thresholds w^{\min} . The threshold value is implied by the number of exceedence values s . The x -axis refers to the proportion of exceedences $s/100,000$ and the y -axis displays t_w . All test statistics are sufficiently negative for variance existence.

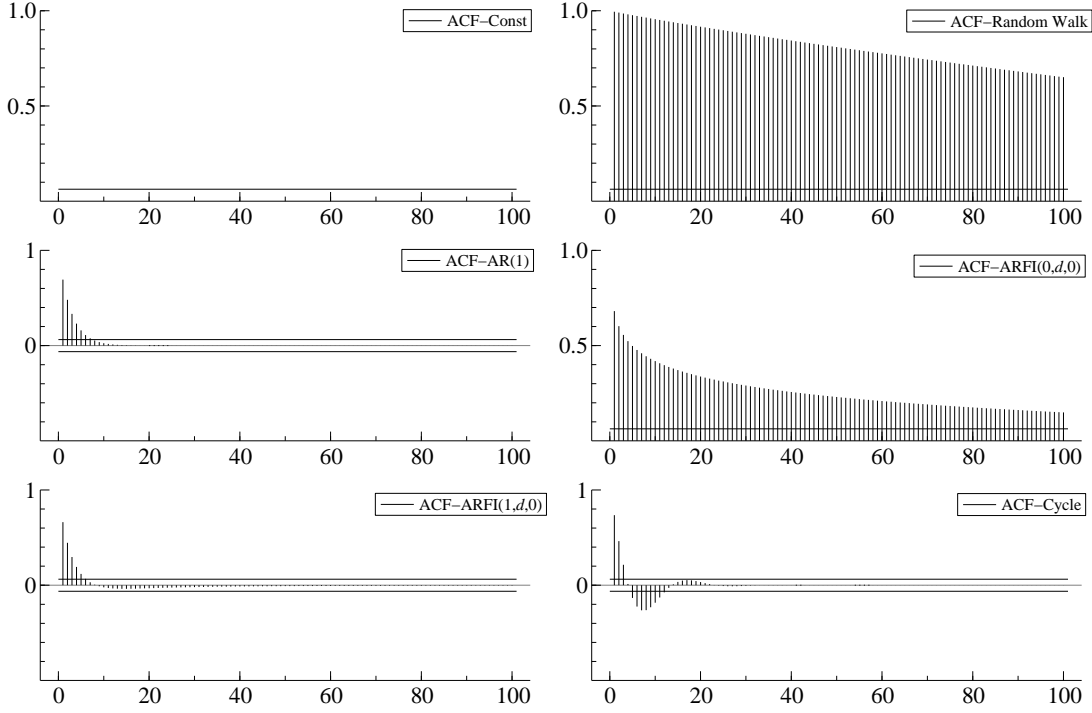


Figure 3: Theoretical autocorrelation functions based on the estimated parameters in the top panel of Table 1 for the stochastic processes u_t .

long-memory process. The autocorrelation function for the stochastic cycle model is first positive for almost 5 lags after which it becomes significantly negative for another 5 lags. This corresponds to the pattern of winning streaks that we observe in the data.

To summarize the in-sample findings. The overall in-sample performance of the models are similar: no large differences in the fit of the models are obtained. The observed average difference in the weights of the two boats is important as a covariate or predictor. At the time that the forecast is usually made, the weights of the boats are known. The models with an autoregressive structure perform relatively better. The out-of-sample forecasting performance is likely to give different results because the implied ACFs are different for different models as indicated in Figure 3.

3 A forty year forecasting assessment

In this section we discuss the out-of-sample forecasting results. All models are used to forecast the Cambridge winning probability repetitively for a period of forty years. The first forecast is made for the race in 1971 and the final forecast is made for the race in 2010. The forecasts are compared across models based on various loss functions and predictive ability tests as discussed by Diebold & Mariano (1995). Further details of our design of the forecasting study are discussed below.

Empirical evidence from out-of-sample forecast performance is generally considered more trustworthy than evidence based on in-sample performances such as the results presented in Section 2. In-sample performance may be more sensitive to data-mining and outliers; see the discussion in White (2000). Also, out-of-sample forecasts provide a more realistic reflection of the real-time information that is available to bookmakers and gambling men.

3.1 Forecasting procedure

We perform an out of sample forecast study where we repetitively forecast the probability that Cambridge wins the Boat Race based on information that is available just before the race takes place. Forecasts for years 1971 until 2010 are computed using a rolling forecast window. The choice for the rolling window is motivated by assumptions underlying the test statistics employed.

To illustrate our procedure in detail, consider the first subsample. It consists of the outcomes corresponding to the years 1829 to 1970. Using this subsample we produce a forecast for 1971. The next subsample contains outcomes from years 1830 to 1971 and produces a forecast for 1972. The final forecast, for 2010, is based on outcomes from 1869 until 2009. In total, we construct $m = 40$ one-year-ahead real-time forecasts for each model.

Let the integer N denote the year 1970, which corresponds to the sample split point and the 114th Boat Race (142 - 28 missing), respectively. The forecasts are computed for $t = N + 1, \dots, N + m$ based on subsamples of the observations y_j, \dots, y_{N+j-1} , for $j = 1, \dots, m$. We estimate the parameter vector ψ for each subsample y_j, \dots, y_{N+j-1} and denote the estimate by $\hat{\psi}_j$. The predicted Cambridge winning probability is computed by

$$\hat{\pi}_{N+j|j,\dots,N+j-1} = \frac{\sum_{i=1}^M \exp(\theta_{N+j|j,\dots,N+j-1}^{(i)}) / (1 + \exp(\theta_{N+j|j,\dots,N+j-1}^{(i)})) w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \hat{\psi}_j)}{\sum_{i=1}^M w(\mathbf{u}^{(i)}, \mathbf{y}; \mathbf{X}, \hat{\psi}_j)}, \quad (13)$$

where

$$\theta_{N+j|j,\dots,N+j-1}^{(i)} = \hat{\mu} + \mathbf{x}_{N+j} \hat{\beta} + u_{N+j|j,\dots,N+j-1}^{(i)},$$

for $j = 1, \dots, m$. The one-year-ahead prediction for the stochastic component $u_{N+j|j,\dots,N+j-1}^{(i)}$ is based on the linear Gaussian importance model $g(\mathbf{y}|\mathbf{u})g(\mathbf{u})$ and is computed by the Kalman filter; see Durbin & Koopman (2012, Chapter 11) for further details.

Let the set of models \mathcal{M} include all models. The models in \mathcal{M} are indexed by k , for $k = 1, \dots, K$, where K is the number of models in set \mathcal{M} . We adjust the notation for the one-year-ahead forecast to reflect that $\hat{\pi}_{N+j|j,\dots,N+j-1}^k$ is the forecast for sample $j = 1, \dots, m$ and model $k = 1, \dots, K$.

3.2 Loss functions

To evaluate the predictive accuracy of the different models, we rely on loss functions. We can refer to these as scores when evaluating probability forecasts. In the literature a large

variety of loss functions exist and, as discussed by Diebold (1993), appropriate loss functions depend on the situation at hand. From a gamblers perspective, the loss function should have the betting odds in each year as an argument. This would allow us to measure the loss in terms of lost investments in a similar way as when trading rules are evaluated; see Hsu & Kuan (2005). However, this information is not available to us.

In our study we rely on more general loss functions which we have computed as follows. For $j = 1, \dots, m$ and $k = 1, \dots, K$, we compute

(i) Brier loss:

$$L_{N+j}^{1,k} = 2(y_{N+j} - \hat{\pi}_{N+j|j,\dots,N+j-1}^k)^2.$$

(ii) Predictive log likelihood loss:

$$L_{N+j}^{2,k} = -y_{N+j} \log(\hat{\pi}_{N+j|j,\dots,N+j-1}^k) - (1 - y_{N+j}) \log(1 - \hat{\pi}_{N+j|j,\dots,N+j-1}^k).$$

(iii) Incorrect loss:

$$L_{N+j}^{3,k} = \begin{cases} 1 & \text{if } \hat{\pi}_{N+j|j,\dots,N+j-1}^k > 0.5 \text{ and } y_{N+j} = 0 \\ 1 & \text{if } \hat{\pi}_{N+j|j,\dots,N+j-1}^k \leq 0.5 \text{ and } y_{N+j} = 1 \\ 0 & \text{else} \end{cases}$$

The Brier loss function, $L_{N+j}^{1,k}$, is proposed by Brier (1950) and is well known for its use in evaluating weather forecasts. It may be viewed as the mean squared error loss function for probabilistic forecasts. The difference is that $\hat{\pi}_{N+j|j,\dots,N+j-1}^k$ is not a forecast for y_{N+j} , but rather a probability statement for the event $y_{N+j} = 1$. The predictive log likelihood loss function, $L_{N+j}^{2,k}$, uses the negative of the log likelihood as a measure of accuracy, which has optimal value zero. The incorrect loss function, $L_{N+j}^{3,k}$, measures the loss arising from event forecasting. It takes values of one or zero, indicating loss or success, respectively.

Using these loss functions we define the relative performance of a model l against another model k , with $k, l \in \mathcal{M}$, by

$$d_j^{s,lk} \equiv L_{N+j}^{s,l} - L_{N+j}^{s,k}, \quad s = 1, 2, 3, \quad l \neq k \quad k, l = 1, \dots, K, \quad j = 1, \dots, m. \quad (14)$$

All predictive ability tests that we consider use functions of the relative performance vectors, $\mathbf{d}^{s,lk} = (d_1^{s,lk}, \dots, d_m^{s,lk})'$, as test statistics. It is easy to show that the vector series $\mathbf{d}^{s,lk}$ is stationary. Let the first moment of the the relative performance indicators be denoted by $\xi^{s,lk} = E[d_j^{s,lk}]$, which we assume independent from j , for all $l, k \in \mathcal{M}$ and $s = 1, 2, 3$. For notational convenience, we drop the dependence on the type of loss function s , when discussing the test statistics below.

3.3 Equal predictive ability

Equal predictive ability (EPA) tests are based upon the null hypothesis that there is no difference in accuracy between the competing models. For comparing models l and k this results in the null hypothesis $H_0 : \xi^{lk} = 0$; see Diebold & Mariano (1995) and West (1996).

The widely used EPA t -type test statistic, first proposed in Diebold & Mariano (1995), is given by

$$T_{lk}^{EPA} \equiv \frac{m^{1/2} \bar{d}^{lk}}{\sqrt{2\pi \hat{f}_{d^{lk}}(0)}}, \quad (15)$$

where $\bar{d}^{lk} = m^{-1} \sum_{j=1}^m d_j^{lk}$ and $\hat{f}_{d^{lk}}(0)$ is a consistent estimate for the spectral density at frequency zero, which may be given by

$$2\pi \hat{f}_{d^{lk}}(0) = \sum_{\tau=-(m-1)}^{m-1} I(\tau/S(m)) \hat{\gamma}^{lk}(\tau), \quad (16)$$

where

$$\hat{\gamma}^{lk}(\tau) \equiv m^{-1} \sum_{s=1}^{m-\tau} (d_s^{lk} - \bar{d}^{lk})(d_{\tau+s}^{lk} - \bar{d}^{lk}), \quad (17)$$

are the sample autocovariance functions and $I(\tau/S(m))$ is the lag window, with truncation lag $S(T)$; see Diebold & Mariano (1995) for further details.

When the models are nested under the null hypothesis of equal predictive ability, the asymptotic theory from West (1996) may not apply. For our models the nesting relationships are non-trivial. It is unclear given the nonlinearities in the observation density, as well as in some of the signal processes, whether the models are nested in the conventional sense. Clark & McCracken (2001) derive asymptotic properties and critical values for the EPA test statistics when models are nested. To avoid further discussions, we view our EPA test statistic as indicative and we assume that it converges to a standard normal distribution as m goes to infinity as shown by Clark & McCracken (2001).

3.4 Results

We discuss the empirical results for our forty years of out-of-sample forecasting and testing. All models from Section 2 with the parameter estimates given in Table 1 are used to obtain forecasts for the Boat Race for the years from 1971 until 2010. We do not use the winner of the toss as a possible predictor as its value was found insignificant for all models. Further, we do not show the results for the ARFI(1, d) model as its results were found indistinguishable from the AR(1) model results. We do include a number of non-parametric forecasting rules for comparison purposes. In particular, we include forecasts that select: last years race winner (the team is in a flow), last years race loser (the team is highly motivated), always Cambridge (duck egg blue supporter) and always Oxford (dark blue supporter). The goal is to show that, at least a selection, of the parametric models of Section 2 are able to outperform

these simple forecasting procedures. The forecasts for the parametric models of Section 2 are computed using a rolling forecasting window. For each model we compute forty forecasts using equation (13) with $M = 1000$ number of draws from the importance density.

Forecasting results

The average loss functions $m^{-1} \sum_{j=1}^m L_{N+j}^{s,k}$ for models $k = 1, \dots, K$ are shown in Table 2. The non-parametric forecasting rules are not evaluated for the log loss function, $s = 2$, as they do not provide probability forecasts.

First we discuss the models that include the average log difference in weight between the boats as a predictor. For all loss functions, the binary model with the stochastic cycle component of Harvey (1989) as the signal u_t provides the lowest loss scores. Also the model that includes the AR(1) component in the signal θ_t yields comparably low scores. The incorrect loss function reveals that our best models (AR(1) and Cycle) are able to predict 30 out of 40 races correctly. This is based on predicting Cambridge as the winner if the predicted probability is larger than one and visa versa. We notice that the forecasting difference between the different parametric signal specification is small. For example, for the Brier loss function, an additional mean squared error of only 0.04 is found between the worst and the best model.

For the models that do not include the difference in the weight of the two boats as a predictor, the differences between the models are larger. This is illustrated in Figure 4, where we present the predicted probability paths for all parametric models. The difference in weight predictor clearly dominates the forecasts when it is included. When it is not included, the constant and random walk models perform much worse, whereas the other parametric models are less affected. Overall we may conclude that the inclusion of the weight predictor is important.

The average loss functions for the non-parametric forecasting rules are presented in the lower panel of Table 2. The Brier loss function attributes very high losses to these models. It is shown that the “always Cambridge” and “loser” rules predict less than 50 % of the races correctly. It seems that always forecasting last years race winner is a reasonable strategy.

Predictive ability results

In Table 3 we present the p -values for the EPA test statistics for the models that include the log difference in weight as a predictor. Each model is compared against all other models for all loss functions. When the p -values are smaller than α , it indicates that the model listed in the row is outperformed by the model listed in the column, with significance level α which we take as $\alpha = 0.1$ in our results. For the Brier loss function, the EPA tests give two main results. First, all non-parametric forecasting rules are significantly outperformed by all parametric models. Especially the rules that predict the “loser” and “always Cambridge” appear to be unsatisfactory when aiming to produce accurate forecasts. Second, the model that includes a random walk process in the signal is outperformed by models that include a more advanced stochastic process in the signal, such as the AR(1), ARFI(0, d) and stochastic cycle processes.

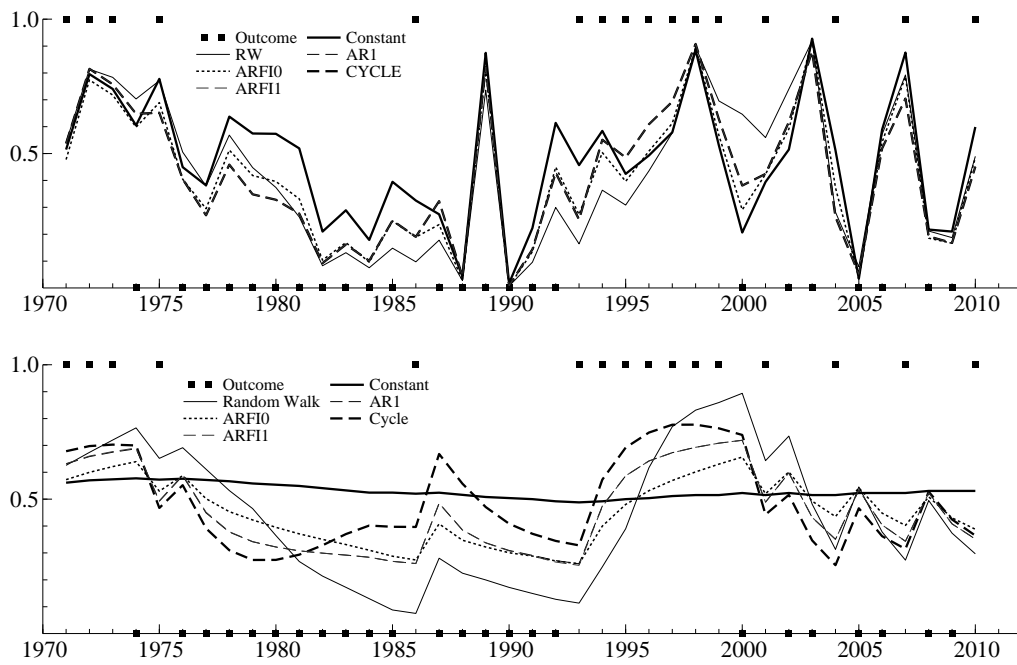


Figure 4: 40 year forecasts for the Boat Race from 1971 until 2010 ($m = 40$ forecasts). Each forecast is based on 142 observations. The top panel shows the forecasts for the parametric models of Section 2 including the weight difference of the two boats as a predictor. The bottom panel shows the forecast for models including only dynamics

Model	Brier	LOG	Incorrect
Including predictors			
Constant	0.419	0.611	0.375
Random walk	0.433	0.625	0.375
AR(1)	0.381	0.561	0.250
ARFI(0,d,0)	0.394	0.578	0.325
CYCLE	0.373	0.555	0.250
Only dynamics			
Constant	0.522	0.715	0.625
Random walk	0.508	0.711	0.425
AR(1)	0.431	0.623	0.350
ARFI(0,d,0)	0.455	0.648	0.375
CYCLE	0.430	0.617	0.350
Non-parametric rules			
Winner	0.700	-	0.350
Loser	1.300	-	0.650
Cambridge	1.200	-	0.600
Oxford	0.800	-	0.400

Table 2: Average loss functions for 40 year forecasting for the Boat Race from 1971 until 2010 ($m = 40$ forecasts). Each forecast is based on 142 observations. Brier corresponds to loss function $m^{-1} \sum_{j=1}^m L_j^{1,k}$, LOG to $m^{-1} \sum_{j=1}^m L_j^{2,k}$ and Incorrect to $m^{-1} \sum_{j=1}^m L_j^{3,k}$. The highlighted numbers indicate the lowest loss per category per loss function.

Perhaps surprisingly, the logistic regression model (constant) is not outperformed by the more advanced dynamic models that take the serial correlation into account. However, the AR(1) and stochastic cycle based models are most unlikely to be outperformed.

The EPA tests based on the log loss function confirm most of the results that were found for the Brier loss function. Again the random walk model is statistically outperformed by the models that include more advanced stochastic processes. The EPA test based on the log loss function is also not able to distinguish between the forecasting performance of the other parametric models. In other words, no evidence is found in favor of either the logistic regression (constant), AR(1), ARFI(0,d), or stochastic cycle models.

The EPA tests based on the incorrect loss function show somewhat different results. The logistic regression (constant) model is now significantly outperformed by the AR(1) and stochastic cycle models. The random walk model is also outperformed by the AR(1) and stochastic cycle models. Interestingly, the forecasting rule that predicts the last years race winner seems to perform much better based on the incorrect loss function. Only the AR(1) and stochastic cycle models are able to reject this model.

In Table 4 we show the results for the models that only depend on dynamics and do not include predictors. Here the differences between the models are more clearly visible. The constant, random walk and all non-parametric forecasting rules are significantly outperformed by the other models for the Brier and log loss functions. For the incorrect loss function the opposite is found. Here only the constant model and the “always Cambridge”

and “loser” prediction rules are outperformed.

When summarizing the testing results for EPA tests, we may conclude that the parametric models predict the Boat Race significantly better when compared to the non-parametric methods based on ad-hoc rules. Furthermore we have provided substantial evidence in favor of models that include an autoregressive component in the signal. These are the models with AR(1) or stochastic cycle process in the signal. There is little evidence against the ARFI(0, d) model but it does not significantly outperform the other models.

4 Conclusion

The forecasting of the outcomes of the yearly Boat Race between Cambridge and Oxford over the forty year period 1971-2010 is extensively evaluated in a real-time study. The accuracy of the forecasts is measured by different loss functions and by equal predictive ability tests. The overall finding is that parametric models predict the outcome of the Boat Race significantly better than ad-hoc methods. Furthermore, models with a latent autoregressive component in the signal produce the most accurate forecasts. Although this study has been mostly fun for us, we do believe that statistical dynamic models have a serious role to play in event forecasting. Since other events may have a more serious impact on us than the Boat Race outcome, the ability to forecasts binary time series accurately is important. The formulation of dynamic models, the development of estimation and forecasting procedures, and the assessment of significant outperformance in forecasting accuracy in the context of binary time series may provide many interesting research questions.

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Equal predictive ability; Including Regressors

Bench. / Alt.	Constant	Random walk	AR(1)	ARFI(0,d)	CYCLE	Winner	Loser	Cambridge	Oxford
Brier loss									
Constant	-	0.624	0.151	0.191	0.138	0.954	1.000	1.000	0.988
Random walk	0.376	-	0.024	0.034	0.086	0.963	1.000	1.000	0.990
AR(1)	0.849	0.976	-	0.812	0.339	0.987	1.000	1.000	0.996
ARFI(0,d)	0.809	0.966	0.188	-	0.250	0.978	1.000	1.000	0.995
CYCLE	0.862	0.914	0.661	0.750	-	0.990	1.000	1.000	0.997
Winner	0.046	0.037	0.013	0.022	0.010	-	0.977	0.984	0.692
Loser	0.000	0.000	0.000	0.000	0.000	0.023	-	0.308	0.016
Cambridge	0.000	0.000	0.000	0.000	0.000	0.016	0.692	-	0.098
Oxford	0.012	0.010	0.004	0.005	0.003	0.308	0.984	0.902	-
Log loss									
Constant	-	0.605	0.119	0.163	0.142	-	-	-	-
Random walk	0.395	-	0.025	0.038	0.113	-	-	-	-
AR(1)	0.881	0.975	-	0.831	0.411	-	-	-	-
ARFI(0,d)	0.837	0.962	0.169	-	0.290	-	-	-	-
CYCLE	0.858	0.887	0.589	0.710	-	-	-	-	-
Winner	-	-	-	-	-	-	-	-	-
Loser	-	-	-	-	-	-	-	-	-
Cambridge	-	-	-	-	-	-	-	-	-
Oxford	-	-	-	-	-	-	-	-	-
Incorrect loss									
Constant	-	0.500	0.042	0.238	0.060	0.409	0.995	0.986	0.586
Random walk	0.500	-	0.024	0.205	0.042	0.391	0.988	0.975	0.596
AR(1)	0.958	0.976	-	0.964	0.500	0.902	1.000	0.999	0.939
ARFI(0,d)	0.762	0.795	0.036	-	0.085	0.609	0.997	0.991	0.782
CYCLE	0.940	0.958	0.500	0.915	-	0.927	0.999	0.999	0.939
Winner	0.591	0.609	0.098	0.391	0.073	-	0.977	0.984	0.692
Loser	0.005	0.012	0.000	0.003	0.001	0.023	-	0.308	0.016
Cambridge	0.014	0.025	0.001	0.009	0.001	0.016	0.692	-	0.098
Oxford	0.414	0.404	0.061	0.218	0.061	0.308	0.984	0.902	-

Table 3: Equal predictive ability tests for comparing 40 year forecasting results for the Boat Race from 1971 until 2010 ($m = 40$ forecasts) for models that include predictors. The t^{EPA} is constructed for different loss functions as discussed in Section 3.3. The highlighted numbers are significant for $\alpha = 0.1$.

Equal predictive ability; Only dynamics

Bench. / Alt.	Constant	Random walk	AR(1)	ARFI(0,d,0)	CYCLE	Winner	Loser	Cambridge	Oxford
Brier loss									
Constant	-	0.407	0.025	0.043	0.024	0.880	1.000	1.000	0.955
Random walk	0.593	-	0.013	0.035	0.064	0.934	1.000	1.000	0.968
AR(1)	0.975	0.987	-	0.943	0.424	0.991	1.000	1.000	0.992
ARFI(0,d,0)	0.957	0.965	0.057	-	0.201	0.974	1.000	1.000	0.989
CYCLE	0.976	0.936	0.576	0.799	-	0.992	1.000	1.000	0.992
Winner	0.120	0.066	0.009	0.026	0.008	-	0.977	0.984	0.692
Loser	0.000	0.000	0.000	0.000	0.000	0.023	-	0.308	0.016
Cambridge	0.000	0.000	0.000	0.000	0.000	0.016	0.692	-	0.098
Oxford	0.045	0.032	0.008	0.011	0.008	0.308	0.984	0.902	-
Log loss									
Constant	-	0.477	0.029	0.050	0.028	-	-	-	-
Random walk	0.523	-	0.014	0.041	0.060	-	-	-	-
AR(1)	0.971	0.986	-	0.940	0.412	-	-	-	-
ARFI(0,d,0)	0.950	0.959	0.060	-	0.201	-	-	-	-
CYCLE	0.972	0.940	0.588	0.799	-	-	-	-	-
Winner	-	-	-	-	-	-	-	-	-
Loser	-	-	-	-	-	-	-	-	-
Cambridge	-	-	-	-	-	-	-	-	-
Oxford	-	-	-	-	-	-	-	-	-
Incorrect loss									
Constant	-	0.018	0.007	0.006	0.007	0.007	0.596	0.327	0.059
Random walk	0.982	-	0.156	0.156	0.201	0.180	0.958	0.942	0.409
AR(1)	0.993	0.844	-	0.673	0.500	0.500	0.980	0.980	0.704
ARFI(0,d,0)	0.994	0.844	0.327	-	0.369	0.352	0.978	0.975	0.596
CYCLE	0.993	0.799	0.500	0.631	-	0.500	0.980	0.984	0.692
Winner	0.993	0.820	0.500	0.648	0.500	-	0.977	0.984	0.692
Loser	0.404	0.042	0.020	0.022	0.020	0.023	-	0.308	0.016
Cambridge	0.673	0.058	0.020	0.025	0.016	0.016	0.692	-	0.098
Oxford	0.941	0.591	0.296	0.404	0.308	0.308	0.984	0.902	-

Table 4: Equal predictive ability tests for comparing 40 year forecasting results for the Boat Race from 1971 until 2010 ($m = 40$ forecasts) for models that only include dynamics. The t^{EPA} is constructed for different loss functions as discussed in Section 3.3. The highlighted numbers are significant for $\alpha = 0.1$.

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